Expected Returns, Yield Spreads, and Asset Pricing Tests

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Abstract

We use information contained in yield spreads to recover investors' *ex ante* required rates of return on corporate securities, and then use these *ex ante* returns to study the pricing of risky assets. Differently from the standard approach, our asset pricing tests do not rely on the use of *ex post* average equity returns as proxies for expected equity returns. We find that: (i) the market beta plays a significant role in the cross-section of expected equity returns, and its role persists even after size and book-to-market factors are accounted for; (ii) the risk premia associated with size and book-to-market are positive, significant, and countercyclical; and (iii) there is little evidence on positive momentum profits. We also find that systematic risk, as captured by common equity factors, is the main driver of the cross-sectional variation in bond yield spreads.

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1 Introduction

The standard asset pricing theory argues that investors demand an *ex ante* premium for acquiring risky securities (Sharpe (1964), Lintner (1965), and Merton (1973)). Since the *ex ante* risk premium is not readily observable, empirical studies commonly use average *realized* returns as a proxy for *expected* equity returns. This practice is justified on grounds that for sufficiently long horizons, the average return will "catch up and match" expected return on equity securities — *ex post* average excess equity returns provide for an easy-to-implement, seemingly unbiased estimate of expected equity risk premium.

Despite its popularity, the above empirical strategy has potentially serious limitations.[1] In particular, the average realized return might not converge to the expected risk premium in finite samples. This, in effect, conditions any inferences based on *ex post* returns on the properties of the particular data under examination.[2] In his AFA presidential address, Elton (1999) observes that there are periods longer than ten years during which stock market realized returns are on average lower than the risk-free rate (1973 to 1984), and periods longer than 50 years in which risky bonds on average underperform the risk-free rate (1927 to 1981). Based on this counter-intuitive evidence, Elton predicts:

> "developing better measures of expected return and alternative ways of testing asset pricing theories that do not require using realized returns have a much higher payoff than any additional development of statistical tests that continue to rely on realized returns as a proxy for expected returns." (p. 1200)

Since most findings in the empirical asset pricing literature were established — and are often revisited — with the use of *ex post* returns, it is natural to ask whether extant inferences about risk–expected return trade-off hold under a direct measure of *ex ante* expected return.

In this paper, we construct an *ex ante* measure of risk premium based on data from bond yield spreads and investigate whether well-known equity factors, such as market, size,

[1] Earlier studies have discussed in some detail the noisy nature of average realized returns in a number of different contexts (see, e.g., Blume and Friend (1973), Sharpe (1978), and Miller and Scholes (1982)).

[2] Complexity in learning about asset pricing formation might also cause *ex post* returns to deviate from their expectations (e.g., Lewellen and Shanken (2002) and Brav and Heaton (2002)).

book-to-market, and momentum, can explain the cross-sectional variation of *expected* (as opposed to average *realized*) stock returns. We also examine whether these equity factors help explain the cross-section of corporate yield spreads.

Our basic approach recognizes that debt and equity are contingent claims written on the same set of real assets, and thus must share the same risk factors that govern the covariance between the underlying firm production process and the aggregate economy. The upshot of recognizing this link is the ability to use corporate bond data to glean additional information about investors' required equity risk premium. In what follows, we derive an analytical formula that links *ex ante* equity risk premia and bond risk premia, after adjusting bond yields for default risk, rating transition risk, and the tax spreads between the corporate and the Treasury bonds.

Why use bond data? While similar information regarding a firm's systematic risk is incorporated both into its equity and bonds, the latter reveal key insights about investors' return expectations. The first thing to notice is that bond yields are calculated in the spirit of internal rates of return. To wit, bond yield is the expected return if the bond does not default and the yield does not change in the subsequent period. Bond prices impound the probability of default, and yield spreads contain the expected risk premium for taking default risk. Controlling for default risk, firms with higher systematic risk will have higher yield spreads; a relationship that holds *period by period*, *cross-sectionally*. This contrasts sharply with what one can learn from equity securities, whose prices reveal little conditional information about expected cash flows and discount rates — one has to rely on a long time series to "back out" the expected return.[3]

Second, and perhaps more importantly, note that the time-variation of expected returns in the equity markets often works *against* the convergence of average realized returns to the expected return. For example, suppose investors require a higher equity risk premium from cyclical firms during economic downturns. To reflect this, those firms' equity prices should drop and their discount rates rise during recessions. Cyclical firms' equity values indeed

[3]As pointed out by Sharpe (1978), the CAPM only holds conditionally and expected return might have nothing to do with future realized returns. Risk premia recovered from bond yields, in contrast, will reflect conditional information.

fall during economic downturns, reflecting value losses in those firms' underlying assets. However, by averaging (*ex post*) a cyclical firm's returns over the course of a recession, one might wrongfully conclude that the cyclical firm is less risky, because of its lower "expected" return. Bond yield spreads, in contrast, increase during recessions: they move in the same direction of the discount rate and spreads are higher for cyclical firms.

Our asset pricing tests provide fresh insights on the determinants of the cross-section of expected returns, complementing the inferences based on the average realized returns. Our main findings are as follows. First, we find that market beta plays a significant role in explaining the cross-sectional variation of expected returns. Importantly, its role persists even after we control for the size and book-to-market factors. This finding is striking given the well-known weak relation between market beta and average returns (e.g., Fama and French (1992)). The contrasts our tests reveal reflect the weak association between our *ex ante* measure of expected return and the commonly used *ex post* measure.

Second, we find that both the expected size premium and the expected value premium are significantly positive throughout our sample period (1973–1998) and are countercyclical. The evidence we present supports the view that size and book-to-market capture relevant dimensions of risk that are priced *ex ante* in equity returns. Further, our finding that the size premium persists over the years contrasts with that of studies that use the average realized returns as the proxy for expected returns.

Third, we find that the expected momentum returns are significantly negative and procyclical. Provided that our measure captures investors' expectations, our tests show that *ex ante* momentum profits do not exist. Instead, our evidence suggests that momentum might be an empirical pattern born out of the use of average realized returns as a proxy for expected returns — momentum might be "more apparent than real" (Schwert (2003)).

Finally, we also examine whether equity factors can explain the cross-section of bond yield spreads. This is important since there are discrepancies in the literature regarding whether the yield spread is related to the equity risk premium. On the one hand, studies on default risk assume that the yield spread is purely idiosyncratic.[4] On the other, empirical

[4]Examples include Bodie et al. (1993), Fons (1994), and Cumby and Evans (1995).

asset pricing papers often use the default spread as an instrument to model the equity risk premium.[5] Sorting out whether the yield spread is mainly driven by idiosyncratic or systematic risk is a relevant matter for empirical research, but surprisingly, little work has been done in this regard. Our focus on equity factors also differentiates us from studies on other yield spread determinants (e.g., Campbell and Taksler (2002) and Chen et al. (2003)).

We find that the Fama-French three-factor model can explain up to 49% of the cross-sectional variation of the yield spreads. The number goes up to 68% when we include other well-known determinants of yield spreads, such as bond rating, maturity, and equity volatility. Thus, for the part of the yield spreads that can be explained, about 72% can be attributed to common equity factors. Our evidence suggests that systematic risk plays a dominant role in driving the cross-section of the yield spreads.

Most empirical studies use the average realized returns as the proxy of expected returns. One important exception that we are aware of is a recent paper by Brav et al. (2003), who following early work by Ang and Petersen (1986), use analyst forecasts to construct an *ex ante* equity risk premium.[6] Our approach differs from that of Brav et al. in that we use bond pricing data, as opposed to analysts' forecasts. In essence, we explore a dimension of investors' information set that is not only different, but uniquely tied to investors' revealed preferences; i.e., comes from their true demand for risky corporate securities. Some of our conclusions regarding the importance of common equity factors resemble those of Brav et al. But importantly, while we find a significantly positive value premium, those authors do not.

The remainder of the paper is organized as follows. Section 2 describes the construction of our *ex ante* equity risk premium measure. Section 3 reports our findings on the time series of common equity factors, the cross-sectional variation of equity risk premium, and that of yield spreads. Section 4 concludes.

[5] Examples include Chen et al. (1986), Keim and Stambaugh (1986), Campbell (1987), Fama and French (1989, 1993), Ammer and Campbell (1993), and Jagannathan and Wang (1996).

[6] Somewhat related to Brav et al. (2003) is Shefrin and Statman (2002), who use ordinal rankings of analyst buy/sell recommendations as a proxy for expected returns.

2 Constructing Expected Equity Returns

We first demonstrate how to recover investors' *ex ante* required return based on information from bonds through a series of propositions (Section 2.1). We then provide the details of the empirical implementation of the method we propose (Section 2.2).

2.1 Methodology

Proposition 1 *Let R^i_{St} be firm i's expected equity return, R^i_{Bt} be its expected debt return, F_{it}, B_{it}, and S_{it} be its assets, debt, and equity values at time t, respectively, and let r_t be the real interest rate. Then:*

$$R^i_{St} - r_t = \left[\left(\frac{\partial S_{it}}{\partial B_{it}}\right)\left(\frac{B_{it}}{S_{it}}\right)\right]\left(R^i_{Bt} - r_t\right). \tag{1}$$

Proof. See Appendix A. ■

Proposition 1 is intuitive. Since both equity and debt are contingent claims written on the same set of productive assets, a firm's equity risk premium is naturally tied to its debt risk premium. Eq. (1) simply formalizes the intuition. The equity risk premium equals the debt risk premium multiplied by two coefficients. The first coefficient is the first derivative of equity with respect to debt, and the second coefficient is the debt-equity ratio.

Empirically, Eq. (1) allows us to recover the expected equity risk premium using bond returns *without* assuming that the average realized equity return is an unbiased measure of expected equity return. The next two propositions introduce our method of constructing expected bond risk premium, R^i_{Bt}–r_t, from observable bond characteristics, without assuming that average realized bond return is an unbiased measure of expected bond return.

Proposition 2 *Let Y_{it} be the yield to maturity, H_{it} be the modified duration, and G_{it} be the convexity of firm i's bond at time t. In the absence of tax differential between corporate bonds and Treasury bonds, the following relationship holds for expected bond excess return and observable bond characteristics:*

$$R^i_{Bt} - r_t = (Y_{it} - r_t) - H_{it}\frac{\mathrm{E}_t\left[dY_{it}\right]}{dt} + \frac{1}{2}G_{it}\frac{\left(dY_{it}\right)^2}{dt} \tag{2}$$

Proof. See Appendix A. ■

Eq. (2) is easy to interpret. The first term in the right hand side is the yield spread between the corporate bond and Treasury bill, which equals the expected excess return of the bond if the bond yield remains constant. The next two terms adjust for the changes in bond yield: the first order change is multiplied by modified duration and the second order change is multiplied by convexity. In essence, Eq. (2) provides a second order approximation of the bond risk premium based on the yield spread.

The next challenge is to model adequately the yield change. The extant literature is rich in models for bond yields.[7] Rather than imposing a parametric model on the yield process, we focus on capturing two important empirical patterns: (i) bond value decreases in the event of default, and (ii) bond ratings generally revert to their long-run means conditional on no-default. Our objective is achieved in the next proposition.

Proposition 3 *Let π_{it} be the expected default probability, dY_{it}^{-} be the yield change conditional on default, and dY_{it}^{+} be the yield change conditional on no-default. Then expected bond excess return is given by:*

$$R_{Bt}^{i} - r_t = (Y_{it} - r_t) + \mathrm{EDL}_{it} + \mathrm{ERND}_{it} \tag{3}$$

where EDL *denotes expected default loss rate, and is defined as:*

$$\mathrm{EDL}_{it} \equiv \pi_{it}\left(-H_{it}\mathrm{E}_t[dY_{it}^{-}] + \frac{1}{2}G_{it}\left(dY_{it}^{-}\right)^2\right)/dt < 0 \tag{4}$$

and ERND *denotes no-default yield change rate, and is defined as:*

$$\mathrm{ERND}_{it} \equiv (1-\pi_{it})\left(-H_{it}\mathrm{E}_t[dY_{it}^{+}] + \frac{1}{2}G_{it}\left(dY_{it}^{+}\right)^2\right)/dt \tag{5}$$

Proof. See Appendix A. ■

Finally, notice that part of the yield spread of corporate bonds over Treasury bonds arises because corporate bond investors have to pay taxes while Treasury bond investors do not.

[7] Structural models of bond yields assume that firms default if their values fall below some boundary. Examples are Merton (1974), Longstaff and Schwartz (1995), Leland (1994, 1998), Leland and Toft (1996), Anderson and Sundaresan (1996), Duffie and Lando (1999), and Collin-Dufresne and Goldstein (2001). See Sundaresan (2001) for a review and Huang and Huang (2003) for an empirical implementation.

Hence, the component in the yield spread that is related to the tax differential should be removed from the spread if one wants to obtain a clean measure of the bond risk premium. Let C_i be the coupon payment for bond i and let τ be the effective tax rate, then:

$$R^i_{Bt} - r_t = (Y_{it} - r_t) + \mathrm{EDL}_{it} + \mathrm{ERND}_{it} - \mathrm{ETC}_{it} \tag{6}$$

where ETC denotes expected tax compensation, and is given by:

$$\mathrm{ETC}_{it} = \left[(1 - \pi_{it})\,\frac{C_i}{B_{it}}\frac{1}{dt} + \mathrm{EDL}_{it}\right]\tau. \tag{7}$$

In Eq. (7), $(1 - \pi_{it})\,\frac{C_i}{B_{it}}\frac{1}{dt}$ is the coupon rate conditional on no-default. The expected default loss rate, EDL, is also included in (7) to capture the tax refund in the event of a default.

Comparison with Merton (1974)

In Merton's (1974) model, equity is an European call option on the underlying asset, and the value of corporate debt, B_{it}, which has face value K and maturity T, is equal to $B_{it} = D_t - P_{it}$, where D_t is price of risk free bond, and P_{it} is a put option.[8] The yield spread can then be calculated as $y_{it} = \log\left(\frac{K}{B_{it}}\right)/T - r$. It is clear that the yield spread is a function of F_{it}/K and firm volatility σ only. It appears as if systematic risk has no effects on the yield spread.

Notice, however, that the firm value process follows $\frac{dF_{it}}{F_{it}} = \mu_i dt + \sigma d\omega_t$, where μ_i is the instantaneous expected return of firm i; determined by its covariance with the stochastic discount factor. Now, consider two firms, 1 and 2, with firm 1 having a higher systematic risk and a higher expected return. Since firm 1's value grows faster than that of firm 2, all else equal, firm 1 will have a lower default probability. Thus, even though both firms have the same yield spread, after adjusting for the fair compensation for default risk, firm 1 has a higher component in the yield spread representing its higher systematic risk. Similarly, if two firms have the same default probability, the firm with higher systematic risk will have higher yield spread.

This example shows that our series of propositions are consistent with Merton's (1974)

[8] Specifically, $P_{it} = Ke^{-rT}N(-d_2) - F_{it}N(-d_1)$ where $d_1 = \frac{\log(F_{it}/K) + \left(r + \frac{1}{2}\sigma^2\right)T}{\sigma\sqrt{T}}$, $d_2 = d_1 - \sigma\sqrt{T}$, K is the face value of debt, F_{it} is the firm value, T is the maturity of debt, σ is firm volatility, and r is the risk free rate.

risk-neutral valuation. In fact, because our formulations are essentially a second order Taylor expansion, the two approaches are mathematically equivalent. In any case, the yield spreads, after properly adjusting for default risk and other components, are fully capable of identifying the cross-sectional variation of systematic risk and expected returns.

2.2 Implementation

This subsection details the empirical operationalization of each of the components of our *ex ante* measure of expected return (Proposition 1): (i) yield spreads ($R^i_{Bt}-r_t$); (ii) expected default loss rates (EDL_{it}); (iii) no-default yields (ERND_{it}); (iv) expected tax compensation (ETC_{it}); and (v) the derivative of equity with respect to debt ($\partial S_{it}/\partial B_{it}$).

Yield Spreads, $R^i_{Bt}-r_t$

We obtain firm-level bond data from the Lehman Brothers fixed income data set, which provides monthly information on corporate bonds, including price, yield, coupon, maturity, modified duration, and convexity, from January 1973 to march 1998. This data set covers a reasonably long period of time and is widely used in empirical research (e.g., Duffee (1999) and Elton et al. (2001)). We only include non-matrix prices because they represent true market quotes. As in Elton et al., we exclude bonds with maturity less than one year since these bonds' prices are less reliable. We include both callable and non-callable bond prices in an effort to retain as many bonds as possible, but our conclusions also hold when only non-callable bonds are used. Finally, we only include bonds issued by non-financial firms.

We obtain Treasury yields for all maturities from the Federal Reserve Board. Following Collin-Dufresne et al. (2001), we compute yield spreads as the corporate bond yields minus the Treasury yields with matching maturities.

Expected Default Loss Rate, EDL_{it}

The expected default loss rate equals the default probability times the actual default loss rate. Moody's publishes information on annual default rates sorted by bond rating from 1970 to 2001, and we use these data to construct expected default probabilities. We note that the literature on default risk typically only uses the unconditional average default probability for each rating and ignores the time variation in expected default probabilities (Elton et al. (2001) and Huang and Huang (2003)). Differently from these papers, our approach is designed to capture time variation in default probability. To do so, we use the three-year moving average default probability from year $t-2$ to t as the one-year expected default probability for year t.[9] For the case of Baa and lower grade bonds, if the expected default probability in a given year is zero, we replace it with the lowest positive expected default probability in the sample for that rating. This ensures that even in occasions of no actual default in three consecutive years, investors still anticipate positive default probabilities.

Table 1 reports the constructed expected default probabilities from 1973 to 1998 based on Moody's data. With only a few exceptions, expected default probabilities decrease with bond ratings. Noteworthy, those default probabilities are typically higher during recessions than during expansions, highlighting the systematic nature of corporate defaults. For example, in the 1990–91 recession, the expected default probability of B3 bonds exceeds 25%, compared to just 5–8% during the late 1990s expansion.

To construct expected default loss rate, EDL_{it}, we still need to gather information on default loss rates. Following Elton et al. (2001), we use the recovery rate estimates provided by Altman and Koshire (1998). Their recovery rates for bonds rated by S&P are: 68.34% (for AAA bonds), 59.59% (AA), 60.63% (A), 49.42% (BBB), 39.05% (BB), 37.54% (B), and 38.02% (CCC). As in Elton et al. (2001), we assume the equivalence between ratings by Moody's and S&P (e.g., Aaa = AAA, ..., Baa = BBB, ..., Caa = CCC), and apply the same recovery rates.

[9] The choice of a three-year window is based on the observation that there are many two-year but few three-year windows without default. While we want to keep the number of years in the window as small as possible, we also want to ensure that expected default probabilities are not literally zero. We have also experimented four other ways to capture the time varying one-year expected default probabilities: (i) using the average one-year default probability from year $t-3$ to $t-1$; (ii) using the actual default probability itself at year t; (iii) using the average default probability from year t to $t+2$; (iv) using the average default probability from year $t+1$ to $t+4$. Results from alternative windows have no bearing on our main conclusions and are available upon request.

Table 1 : Three-Year Moving Average Annual Default Probability (in Percent)

This table reports the three-year moving average annual default rates for corporate bonds categorized by ratings, where the three-year window includes the current year and the previous two years. The table is constructed using annual default rate data from Moody's.

Year	Aaa	Aa1	Aa2	Aa3	A1	A2	A3	Baa1	Baa2	Baa3	Ba1	Ba2	Ba3	B1	B2	B3
1973	0	0	0	0	0	0	0	0.343	0.343	0.343	0.430	0.430	0.430	4.920	4.920	4.920
1974	0	0	0	0	0	0	0	0.343	0.343	0.343	0.430	0.430	0.430	5.937	5.937	5.937
1975	0	0	0	0	0	0	0	0.343	0.343	0.343	0.633	0.633	0.633	5.547	5.547	5.547
1976	0	0	0	0	0	0	0	0.280	0.280	0.280	0.833	0.833	0.833	4.633	4.790	5.093
1977	0	0	0	0	0	0	0	0.280	0.280	0.280	0.867	0.867	0.867	3.427	3.583	3.887
1978	0	0	0	0	0	0	0	0.280	0.280	0.280	0.887	0.887	0.887	3.240	3.397	3.700
1979	0	0	0	0	0	0	0	0.280	0.280	0.280	0.707	0.707	0.707	3.240	3.397	3.700
1980	0	0	0	0	0	0	0	0.280	0.280	0.280	0.673	0.673	0.673	3.833	3.990	4.293
1981	0	0	0	0	0	0	0	0.280	0.280	0.280	0.450	0.450	0.450	3.527	3.683	3.987
1982	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.290	1.213	1.213	1.213	3.987	3.987	3.987
1983	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.290	1.213	1.213	1.940	2.643	5.633	8.270
1984	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.550	1.457	1.607	1.940	3.093	10.387	7.740
1985	0	0	0	0	0	0	0	0.280	0.280	0.540	0.673	1.223	2.270	3.750	12.053	11.557
1986	0	0	0	0	0	0	0	0.280	0.280	2.053	0.823	1.480	2.547	5.943	14.277	10.943
1987	0	0	0	0	0	0	0	0.280	0.280	1.793	1.680	1.260	3.387	5.640	9.460	13.433
1988	0	0	0	0	0	0	0	0.280	0.280	1.793	1.680	0.860	2.993	5.627	9.290	12.053
1989	0	0	0	0.467	0	0	0	0.280	0.453	0.543	1.650	1.067	3.417	5.170	6.493	13.213
1990	0	0	0	0.467	0	0	0	0.280	0.453	0.543	1.297	1.690	3.740	6.390	12.423	19.400
1991	0	0	0	0.467	0	0	0	0.440	0.453	0.543	1.507	1.690	6.173	6.957	14.370	25.633
1992	0	0	0	0	0	0	0	0.440	0.280	0.280	1.387	1.227	4.850	5.220	12.123	27.297
1993	0	0	0	0	0	0	0	0.440	0.280	0.280	0.767	0.430	3.793	3.463	6.413	21.480
1994	0	0	0	0	0	0	0	0.280	0.280	0.280	0.557	0.430	0.693	2.083	3.387	14.690
1995	0	0	0	0	0	0	0	0.280	0.280	0.280	0.557	0.430	1.020	3.190	4.993	7.877
1996	0	0	0	0	0	0	0	0.280	0.280	0.280	0.430	0.430	0.913	2.473	3.840	5.170
1997	0	0	0	0	0	0	0	0.280	0.280	0.280	0.430	0.430	0.873	2.183	3.120	4.957
1998	0	0	0	0	0	0	0	0.280	0.293	0.280	0.430	0.490	0.663	1.443	3.523	5.460

No-Default Yield Change Rate, ERND_{it}

To calculate the expected return due to yield change conditional on no-default, ERND_{it}, we need to calculate dY_{it}^{+}, the yield change conditional on no-default.

Evidence on Mean-reverting Default Probabilities Empirically, if a bond does not default, its default probability typically mean-reverts. In Table 2 we report conditional default probabilities from one to 20 years conditional on no-default in the previous year. These probabilities are constructed by using the one-year default transition matrices provided by Moody's and S&P Corporation.

Table 2 : Annual Default Probability Conditional on No-Default in the Previous Year (in Percent)

This table reports the annual default probability conditional on no-default in the previous year. The table is constructed using the average one-year rating transition matrix of Moody's and that of S&P Corporation, reported in Table V of Elton et al. (2001).

Year	Aaa	Aa	A	Baa	Ba	B	Caa
1	0	0	0.052	0.158	1.402	7.403	22.289
2	0.001	0.011	0.094	0.312	1.949	7.459	19.278
3	0.004	0.024	0.139	0.467	2.348	7.337	16.427
4	0.008	0.042	0.188	0.615	2.633	7.112	13.887
5	0.013	0.062	0.239	0.752	2.832	6.830	11.726
6	0.020	0.084	0.290	0.877	2.963	6.521	9.947
7	0.029	0.109	0.343	0.987	3.044	6.204	8.511
8	0.039	0.136	0.397	1.085	3.084	5.889	7.368
9	0.051	0.165	0.449	1.169	3.094	5.585	6.461
10	0.065	0.195	0.500	1.243	3.081	5.295	5.742
11	0.080	0.226	0.550	1.304	3.051	5.022	5.169
12	0.096	0.259	0.597	1.356	3.009	4.767	4.707
13	0.114	0.291	0.643	1.399	2.957	4.528	4.330
14	0.133	0.324	0.686	1.435	2.898	4.306	4.019
15	0.153	0.358	0.727	1.463	2.837	4.100	3.758
16	0.175	0.391	0.765	1.486	2.771	3.910	3.535
17	0.197	0.425	0.802	1.503	2.706	3.733	3.344
18	0.220	0.458	0.835	1.516	2.639	3.569	3.175
19	0.243	0.490	0.867	1.525	2.574	3.417	3.027
20	0.268	0.522	0.895	1.530	2.510	3.276	2.894

It is clear from Table 2 that, conditional on no-default, annual default probabilities increase over the years for bonds with originally high rating, but they decrease for bonds with originally low rating. For example, at year one, the one-year ahead default probability for Caa bonds is 22.28%. The one-year default probability then goes down to 19.28% in the second year and to 16.43% in the third year. Since mean-revering default probabilities imply mean-reverting yields, high-quality bonds can have positive credit spreads even though their one-year default rates are close to zero.

Further evidence on the mean reversion of yield spreads is provided in Table 3. On an annual basis, we pool together all bonds belonging to the same Moody's rating category in the Lehman Brothers data set and study the changes in cumulative average ratings and yield spreads over the following three years. We assign numeric numbers, from one to seven, to

Table 3 : Evolution of Ratings and Yield Spreads in Corporate Bonds

In this table we use Lehman Brothers Fixed Income dataset (January 1973 to March 1998) to form cohorts of bonds with the same initial rating for each year. We then report the average rating and yield spread changes for the same initial rating groups. Ratings from Aaa to Caa are assigned integer numbers from one to seven, with higher numbers indicating lower ratings. Changes in yield spreads are in percent. The t-statistics for the changes are reported in parentheses.

Year	Changes in	Aaa	Aa	A	Baa	Ba	B	Caa
1	Rating	0.080	0.065	0.025	0.013	-0.042	-0.036	-0.121
		(4.99)	(10.69)	(9.14)	(2.30)	(3.30)	(5.70)	(3.18)
	Yield Spread	0.032	0.058	0.073	0.093	0.002	0.438	-1.272
		(1.39)	(5.06)	(9.30)	(5.67)	(0.04)	(4.92)	(1.72)
2	Rating	0.232	0.174	0.064	0.029	-0.089	-0.070	-0.360
		(7.93)	(15.41)	(11.50)	(2.33)	(3.35)	(5.28)	(3.71)
	Yield Spread	0.074	0.123	0.115	0.115	0.071	0.860	-1.433
		(2.11)	(6.52)	(9.26)	(3.84)	(0.88)	(5.21)	(1.51)
3	Rating	0.434	0.287	0.114	0.047	-0.148	-0.102	-0.649
		(10.10)	(17.14)	(13.24)	(2.55)	(3.96)	(4.57)	(5.72)
	Yield Spread	0.066	0.169	0.138	0.138	-0.116	1.286	-2.423
		(1.46)	(7.03)	(8.28)	(3.00)	(1.32)	(5.25)	(2.83)

bonds rated from Aaa to Caa, with a lower number corresponding to a better rating.

Table 3 shows that the ratings of high-quality bonds (Aaa, Aa) indeed decline over time while their yield spreads increase. For example, the rating of Aa-rated bonds, conditional on no-default, increases by 0.065 after a year, where an increase of one indicates a full downgrade to grade A. Accordingly, the average yield spread of Aa bonds increases by 5.8 basis points. In contrast, the ratings of low-quality bonds (Caa) improve over time and their yield spreads decline. The evidence clearly shows mean reversion in yield spread conditional on no-default.

Modeling Mean-reverting Yield Changes We adopt the following three-step procedure to recover dY_{it}^{+}, the yield change conditional on no-default, from the data. First, we construct the cumulative default probability for each maturity using Table 2. For example, the conditional default probabilities for a bond initially rated Baa are 0.16% and 0.31% for the first two years, respectively. Assuming that the default rate is the same within a given year, the cumulative default probabilities are 0.16%, 0.16%, 0.47% ($= 0.16\% + (1 - 0.16\%) \times 0.31\%$), and again 0.47% for 0.5-year, 1-year, 1.5-year, and 2-year maturity, respectively.

Second, for each bond we calculate the expected cash flow, while taking into account possible default. The expected cash flow for a particular coupon date before maturity is equal to: coupon payment × [1 − cumulative default probability × (1 − recovery rate)], where the recovery rates are from Altman and Koshire (1998). We calculate the present value of the bond by discounting its expected cash flows by the corresponding Treasury yields with matching maturities.[10] After we obtain bond prices, we then calculate bond yields.

Accordingly, suppose the bond in the previous numerical example has two years to maturity and the coupon rate is 8% with face value of $100. Further assume that the current Treasury yield, with annualized semi-annual compounding, is 8% for two-year maturity. Without default, the cash flows for the bond are $4, $4, $4, and $104 for the four half-year periods. The recovery rate from Altman and Koshire (1998) for the Baa-rated bond is 49.42%. With default risk, the expected cash flows are (1 − 0.16% × (1 − 49.42%)) × 4, (1–0.16%×(1–49.42%))×4, (1 − 0.4695% × (1–49.42%))×4, and (1–0.4695%×(1–49.42%))×104, respectively. The present value, when we use the discount rate of 8%, is thus $99.77. With the promised cash flows of $4, $4, $4, and $104, and the price at $99.77, the bond yield equals 8.12%.

Now, assume that the bond does not default within the first year. Conditional on that event, the bond maturity decreases by one year, and the second-year conditional default probability reported in Table 2 becomes the first-year default probability for this "new" bond. One can iterate over the last two steps to calculate the price and yield for the new bond. Because conditional default probabilities of high grade bonds will increase in the second year, bond prices will decrease and yields will increase, revealing a downgrading trend. Similarly, because conditional default probabilities will decrease for low grade bonds in the second year, bond prices will increase and yields will decrease, representing an upgrading trend. The yield difference between the last two steps will be our proxy for the yield change conditional on no-default within the first year. As expected, this yield change, dY_{it}^{+}, will be positive for high grade bonds but negative for low grade bonds.

Let's again consider our numerical example. After one year, conditional on no-default, the

[10]This is equivalent to calculating the fair price of the bond by a risk-neutral investor.

new cumulative default rates will be 0.31% and 0.31% for the 0.5-year and 1-year maturities. Using our method to calculate the expected cash flows for this bond, we find the new price to be \$99.85 and the yield to be 8.17%. Thus, the bond yield will go up by five basis points due to the expected increase of default probability. The five basis points will be used as dY_{it}^{+} in calculating ERND_{it}, the expected return due to yield change conditional on no-default.

In sum, ERND_{it} is a function of rating-specific default probability, bond specific maturity, duration, convexity, and the Treasury yield for a given month. Although tedious, our method ensures that this component of credit spread dynamics is captured with the best available information for the particular bond at any given time.

Expected Tax Compensation, ETC_{it}

To calculate the expected tax compensation given by (7), we follow Elton et al. (2001) and set the effective tax rate to be 4% for all bonds. This completes our construction of the four components in the bond risk premium formula (6).

Constructing Expected Equity Returns

Armed with a measure of expected bond risk premium, it is straightforward to use Proposition 1 to obtain expected equity risk premium. Note, though, that we don't directly observe $\partial S_{it}/\partial B_{it}$. This derivative needs to be estimated from the data and the following steps describe our estimation procedure.

First, we combine the bond data with CRSP monthly data to obtain market capitalization for equity, and then merge it further with Compustat to gather information on firm leverage. The final merged Lehman Brothers/CRSP/Compustat data set includes 1023 non-financial firms covering the period from January 1973 to March 1998.

Second, for each firm, we calculate the change of S_{it} as the market capitalization change for each month. We also add together the value change for each bond within each firm. Our bond data covers approximately 50% of all debt for each Compustat-matched firm, and in order to obtain the change of B_{it}, we multiply the total bond value change by the ratio of book debt from Compustat to the total bond face value from our Lehman Brothers data set.

We then obtain (projected) estimates for $\partial S_{it}/\partial B_{it}$ via a pooled panel regression of $\partial S_{it}/\partial B_{it}$ on a constant and the debt/equity ratio (t-statistics in parentheses):

$$\frac{\partial S_{it}}{\partial B_{it}} = \underset{(49.50)}{12.83} - \underset{(12.55)}{0.58}\,\frac{B_{it}}{S_{it}}. \tag{8}$$

The predicted negative relation between $\partial S_{it}/\partial B_{it}$ and B_{it}/S_{it} agrees with theoretical priors. To see this, consider the simple model of Merton (1974), in which the equity S is a European call option on the underlying asset F. The debt is a zero coupon bond with face value of K and maturity of T years. The Black and Scholes' (1973) formula implies that: $\frac{\partial S}{\partial B} = \frac{N(d_1)}{1-N(d_1)}$, where $N(\cdot)$ is the cumulative density function of normal distribution, $d_1 \equiv \frac{\log(F/K)+\left(r+\frac{1}{2}\sigma^2\right)T}{\sigma\sqrt{T}}$, and σ is the volatility of the firm's asset return. Because d_1 is always positive, $N(d_1) > 0.5$ and $\frac{N(d_1)}{1-N(d_1)} > 1$. Moreover, because d_1 decreases with K/F, $\frac{N(d_1)}{1-N(d_1)}$ decreases with the leverage ratio. That is, for firms with higher leverage, a unit change of debt value is associated with a smaller change in the equity value. Intuitively, given one unit of firm value change, the change in equity value will be smaller if the debt portion is larger.[11]

Having developed empirical counterparts for each of the components of our *ex ante* return measure, we substitute those estimates into Eq. (1) for each firm-month in our sample. We can now study the pricing of risky securities with a direct measure of expected returns.

3 Findings

We first report summary statistics of yield spreads and expected bond risk premia that are used in our expected equity return measure (in Section 3.1). Section 3.2 studies the properties of the common equity factors of Fama and French (1993) and momentum under our *ex ante* measure. Section 3.3 examines the cross-sectional variation of the expected equity risk premium. Section 3.4 shows the robustness of the results derived from our *ex ante* approach. In Section 3.5, we look at the cross-sectional variation of yield spreads.

[11]The specification in Eq. (8) is admittedly simple. We conduct extensive robustness checks below (Section 3.4) using alternative specifications for $\partial S_{it}/\partial B_{it}$.

Table 4 : Summary Statistics of Yield Spreads and Expected Bond Risk Premium By Bond Ratings

This table reports summary statistics, including mean, standard deviation (std), min, max, and autocorrelations of orders 1 (ρ_1), 2 (ρ_2), 6 (ρ_6), and 12 (ρ_{12}), of yield spread (Panel A) and expected bond risk premium (Panel B) for bonds rated from B to Aaa. The mean, min, and max are in annualized percent.

	Panel A: Yield Spread							
Rating	mean	std	min	max	ρ_1	ρ_2	ρ_6	ρ_{12}
Aaa	0.850	0.049	0.302	3.092	0.948	0.909	0.781	0.540
Aa	0.897	0.053	0.419	2.286	0.965	0.930	0.800	0.616
A	1.093	0.063	0.608	2.514	0.948	0.902	0.691	0.409
Baa	1.805	0.106	0.456	4.310	0.955	0.910	0.708	0.445
Ba	2.967	0.175	1.738	6.785	0.852	0.793	0.626	0.419
B	5.494	0.324	2.671	18.460	0.957	0.932	0.801	0.546
	Panel B: Expected Bond Risk Premium							
Rating	mean	std	min	max	ρ_1	ρ_2	ρ_6	ρ_{12}
Aaa	0.464	0.027	-0.027	2.653	0.942	0.897	0.748	0.470
Aa	0.465	0.027	0.011	1.863	0.960	0.919	0.760	0.543
A	0.607	0.035	0.133	2.045	0.943	0.891	0.655	0.351
Baa	0.930	0.054	-0.017	3.379	0.946	0.892	0.666	0.408
Ba	1.053	0.062	0.110	2.917	0.816	0.722	0.432	0.107
B	2.238	0.132	0.066	10.980	0.844	0.765	0.596	0.311

3.1 Yield Spreads and Expected Bond Risk Premium

Table 4 reports summary statistics of yield spreads and constructed bond risk premia for B- through Aaa-rated bonds. Because data are not available on time-varying default rates for bonds rated Caa or lower, we delete these bonds from the sample. (These bonds consist of about one percent of all bonds.) We construct firm-level bond risk premium as the simple average of the risk premia of all the bonds issued by the firm, but emphasize that the use of value-weighted averages yields very similar results, as different bonds issued by the same firm earn very similar risk premia.

Table 4 shows that the yield spread and the expected bond risk premium increase as the bond rating decreases. The bond risk premium for Aaa-rated bonds is on average 0.46% per annum, and it goes up to 2.24% for B-rated bonds. This evidence suggests that lower graded bonds are systematically riskier than higher graded bonds. Both the yield spread and the expected bond risk premium are highly persistent. The first-order autocorrelations range from 0.82 to 0.97, and the 12^{th}-order autocorrelations range from 0.10 to 0.60.

3.2 Common Factors in Expected Equity Returns

We define the market equity risk premium as the value-weighted average equity risk premia of all firms. The expected returns of other common factors, including size and book-to-market, are constructed following Fama and French (1993).[12] At each month, we also sort stocks on the basis of their realized equity return in the past 12 months into winners (W>70%), medium (70%≥M≥30%), and losers (L<30%) categories.[13] The momentum factor is calculated as the winner-minus-loser (WML) portfolio.

Descriptive Statistics

Panel A of Table 5 reports the *ex ante* return summary statistics of the four common equity factors we consider. The expected market risk premium is on average 3.93% per annum, with a low standard deviation of 0.15%. The expected size premium is on average 5.68% (standard deviation of 0.27%) and the expected value premium is on average 9.04% (standard deviation of 0.33%), both of which are highly significant. The momentum factor, in contrast, earns a significantly negative expected return of -2.02% (standard deviation of 0.29%). This evidence suggests that the momentum anomaly (e.g., Jegadeesh and Titman (1993)) might arise from the use of average realized returns as a "poor" proxy for expected returns.

Panel B of Table 5 reports the *ex ante* return correlation matrix of the four equity factors. The expected return of the market factor is positively correlated with the size factor (0.19) and the book-to-market factor (0.68), but is negatively correlated with the momentum factor (-0.44). We also find that the market portfolio of the bond earns on average 0.42% per annum (standard deviation of 0.02%). The equity market risk premium is highly positively

[12]In June of every year, we sort firms according to the current month's equity market capitalization into big and small categories using the 50-50% cutoff points. We sort firms on their book-to-market ratios into High, Median, and Low categories using the 30-40-30% cutoff points. The size factor and the book-to-market factor are mutually stratified. The size premium is calculated as (SH+SM+SL−BH−BM−BL)/3, where SH represents the weighted average equity return for all firms that belong to the small S and high book-to-market H categories, and other portfolios are defined similarly. The value premium is calculated as (SH+BH−SL−BL)/2.

[13]We skip one month to avoid market microstructure difficulties and hold the portfolios for 12 months.

Table 5 : Summary Statistics of Expected Returns of Common Equity Factors

This table reports summary statistics of expected returns of common equity factors, including market excess return (MKT), SMB, HML, and WML (the momentum factor). Panel A reports mean, standard deviation (std), min, max, and autocorrelations of orders 1 (ρ_1), 2 (ρ_2), 6 (ρ_6), and 12 (ρ_{12}). Panel B reports the results of market regressions for SMB, HML, and WML, including the intercepts (α) and the slopes (β) as well as their Newey-West t-statistics. Finally, Panel C reports the correlation matrix for these four factors. The numbers of mean, min, max, and α are in annualized percent. All cross-correlations in Panel C are significant at the 1-percent test level or lower. All the t-statistics are adjusted for heteroscedasticity and autocorrelations of up to 12 lags.

	Panel A: Summary Statistics							
	mean	std	min	max	ρ_1	ρ_2	ρ_6	ρ_{12}
MKT	3.926	0.149	0.220	15.549	0.888	0.848	0.698	0.408
SMB	5.677	0.270	-2.749	22.774	0.921	0.875	0.642	0.353
HML	9.042	0.332	-3.225	26.314	0.896	0.837	0.701	0.554
WML	-2.021	0.291	-18.177	12.513	0.834	0.768	0.350	0.159
	Panel B: Cross-Correlations				Panel C: Market Regressions			
	MKT	SMB	HML	WML	α	t_α	β	t_β
MKT	1	0.191	0.677	-0.439	na	na	na	na
SMB		1	-0.202	-0.209	4.410	9.084	0.327	3.119
HML			1	-0.274	3.498	7.810	1.430	14.784
WML				1	1.265	2.642	-0.860	-8.201

correlated with the market bond risk premium; with a correlation coefficient of 0.86. This is not surprising since, as contingent claims on the same productive assets, equity and bond should share similar risk factors.

Panel C reports the market regressions of SMB, HML, and WML. The unconditional alphas of SMB and HML are 4.41% and 3.50% per annum, and are highly significant (t-statistics of 9.08 and 7.81, respectively). The unconditional betas of SMB and HML are also positive and statistically significant. WML has a positive unconditional alpha of 1.27% (t-statistic of 2.64), but a negative unconditional beta of -0.86 (t-statistic of -8.20).

Business Cycle Properties

We now investigate the cyclical properties of the expected returns for the four equity factors during the 1973–1998 period. Following the empirical business cycle literature (e.g., Stock and Watson (1999, Table 2)), Table 6 reports the cross correlations (with different leads and

Table 6 : Cross Correlations with Cyclical Component of Industrial Production

This table reports the cross correlations of expected returns of equity factors with the cyclical component of the industrial production index, corr(r_t, y_{t+k}), for different leads and lags, k. The cyclical component of the real industrial production index (obtained from FRED) is estimated by passing the raw series through the Hodrick and Prescott (1997) filter. p-values of the cross correlations are reported in parentheses.

r_t	corr(r_t, y_{t+k})												
	-24	-12	-6	-3	-2	-1	0	1	2	3	6	12	24
MKT	0.222	0.193	0.070	-0.051	-0.122	-0.195	-0.241	-0.289	-0.314	-0.313	-0.320	-0.198	0.228
	(0.00)	(0.00)	(0.23)	(0.38)	(0.04)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
SMB	-0.136	0.171	0.056	-0.128	-0.199	-0.275	-0.357	-0.426	-0.463	-0.473	-0.425	-0.086	0.373
	(0.03)	(0.01)	(0.37)	(0.04)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.18)	(0.00)
HML	0.145	0.049	-0.044	-0.162	-0.222	-0.269	-0.295	-0.309	-0.308	-0.299	-0.227	-0.117	0.154
	(0.02)	(0.43)	(0.48)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.07)	(0.02)
WML	-0.209	-0.113	0.099	0.224	0.249	0.268	0.273	0.267	0.254	0.259	0.252	0.062	-0.082
	(0.00)	(0.06)	(0.10)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.31)	(0.19)

lags) of the expected returns with the cyclical component of the real industrial production index. The industrial production index is obtained from the monthly database of the Federal Reserve Bank of Saint Louis. We follow Stock and Watson (1999) by removing from the output series its long-run growth component as well as those fluctuations that occur over periods shorter than a business cycle, which arise from temporary factors such as measurement errors. This is achieved by passing the raw industrial production index through the Hodrick and Prescott (1997) filter. Finally, following Hodrick and Zhang (2001), we set the monthly smooth parameter in the Hodrick-Prescott filter to be 6,400.

Table 6 reports a number of interesting patterns. First, the expected market risk premium is negatively correlated with the cyclical component of output. The cross correlations are mostly negative and significant across different leads and lags. This evidence suggests that the expected market risk premium is countercyclical; i.e., investors demand a higher risk premium in recessions. This finding speaks to the criticism voiced by Elton (1999) that *ex post* equity returns go *down* in recessions and thus fail to capture investors' (presumably) heightened required returns from risky assets in uncertain environments.

Second, both the expected size premium and the expected value premium are negatively correlated with the cyclical component of output. In other words, investors seem to perceive small and value stocks as riskier securities than big and growth stocks *ex ante*, charging a countercyclical risk premium for holding those assets. Note that our evidence on the countercyclical expected value premium is consistent with the recent theoretical work of Zhang (2003). The final noticeable feature of Table 6 concerns the cyclical properties of momentum. In contrast to other equity factors, expected momentum is strongly *procyclical*; the cross correlations between expected momentum returns and output are positive and significant at most leads and lags.

Average Realized Excess Returns Versus Expected Risk Premia

Since we propose that our proxy of expected equity returns might provide for additional insights into the pricing of equity securities, it is important to show how different our *ex ante* proxy is from the one commonly used in the literature. To this end, we perform predictive regressions of future, realized equity factor returns on *ex ante* expected returns.

Table 7 reports the results. Four different horizons are considered: six-month, 12-month, 24-month, and 36-month. Panel A shows that our proxy of expected market risk premium is close to the average market excess return in short horizons up to 12 months. The slope coefficient is close to 0.5 in the six-month regression and close to one in the 12-month regression. However, our proxy diverges from average realized returns over longer horizons; the slopes associated with the 24- and 36-month regressions are not statistically distinguishable from zero.

Panel B reports a more drastic divergence between our proxy of expected return and the average realized return. The slopes of regressing the realized SMB returns on its expected returns across all of the horizons are *negative*! This suggests an explain for why we document a significantly positive size *ex ante* premium over the entire 1973-1998 period, while most studies using average realized returns report a negative size premium in comparable sample periods. Based on the evidence that the size premium disappears after the publication of its discovery in Banz (1981), Schwert (2003) argues that academic research has made the market more efficient. Our evidence, nonetheless, suggests that the disappearance (or instability) of size factor could stem from the use of average realized return as a proxy for expected returns.

Table 7 : Regressing Realized Equity Factor Returns Onto Their Constructed Expected Returns

This table reports predictive regressions of future, realized equity factor returns including market excess return (Panel A), SMB (Panel B), HML (Panel C), and WML (Panel D) on their respective expected returns constructed using the methods discussed in Section 2.1. Four different predictive horizons are considered: (i) six-month; (ii) 12-month; (iii) 24-month; and (iv) 36-month horizons. All the t-statistics are adjusted for heteroscedasticity and autocorrelations of up to 12 lags.

Panel A: MKT					Panel B: SMB				
	6-month	12-month	24-month	36-month		6-month	12-month	24-month	36-month
coef	0.508	0.838	0.050	0.674	coef	-0.024	-0.139	-0.312	-0.274
t-stat	2.047	2.610	0.125	1.505	t-stat	-0.298	-1.350	-2.339	-1.897
R^2	0.014	0.023	0.000	0.009	R^2	0.000	0.007	0.023	0.016
Panel C: HML					Panel D: WML				
	6-month	12-month	24-month	36-month		6-month	12-month	24-month	36-month
coef	0.100	0.169	0.498	0.406	coef	0.634	0.825	0.930	0.474
t-stat	1.069	1.358	2.604	1.731	t-stat	5.409	5.394	5.481	3.150
R^2	0.005	0.007	0.028	0.013	R^2	0.096	0.097	0.104	0.039

3.3 The Cross-Section of Expected Equity Returns

We now examine the cross-sectional variation of *ex ante* equity returns. Using the Fama and MacBeth (1973) methodology, we study whether the loadings on common equity factors, including the market beta, size, book-to-market, and momentum, have power in explaining the cross-section of expected returns. Both covariances and characteristics are used in our asset pricing tests.

Covariance-Based Tests

Our covariance-based tests are conducted in two steps. In the first step, for each individual stock and month, we run the time series regression of the equity returns in the past 60 months (with at least 24 months of data available):

$$r^i_{t+1} = \alpha_i + \beta^j_{\mathrm{MKT}}\mathrm{MKT}_{t+1} + \beta^i_{\mathrm{SMB}}\mathrm{SMB}_{t+1} + \beta^i_{\mathrm{HML}}\mathrm{HML}_{t+1} + \beta^i_{\mathrm{WML}}\mathrm{WML}_{t+1} + \varepsilon_{it+1}. \quad (9)$$

r^i_{t+1} is the realized excess equity return of stock i from time t to $t+1$ over the one-month Treasury bill rate, and MKT, SMB, HML, and WML are the excess return factors of market, size, book-to-market, and momentum taken from Kenneth French's website.

In the second step, we run cross-sectional regressions, month by month, of firm-specific, *expected* equity returns on the factor loadings estimated from Eq. (9). The time series average of the coefficients are regarded as the risk premia associated with the loadings. To adjust the standard errors of the coefficients for their persistence (high autocorrelations), we follow Pontiff (1996) by regressing the slopes on an intercept term and modeling the residuals as a 12-order autoregressive process. The standard errors of the intercepts are then used as the corrected standard errors in computing Fama-MacBeth t-statistics.

Panel A of Table 8 reports the cross-sectional regressions with the factor loadings. The column denoted Model 1 shows that when the equity beta alone is used, the slope coefficient is a positive 2.86%; significant at the 1% test level. In other words, stocks with higher loadings on the market factor will have higher expected excess return, consistent with standard asset pricing models. The column denoted Model 2 indicates that, when the loadings on SMB and HML are also included, all three Fama-French factors have positive coefficients that are significant. This suggests that the market beta is priced even in the presence of SMB and HML factor loadings, and that the Fama-French (1993, 1996) factors are *ex ante* priced. Finally, Model 3 shows that the Fama-French three factors remain positive and significant even after including the loadings on the momentum factor in the regression. The premium associated with the momentum factor loading, however, is negative and insignificant.

Characteristic-Based Tests

In Panel B of Table 8, we retain the loading on the market factor, but replace the other loadings with firm characteristics. That is, we use the natural logarithm of size, the book-to-market ratio, and the prior equity return to replace their corresponding factor loadings. Model 4 shows that the market beta is priced and its slope coefficient is significant for the characteristics of size and book-to-market in the cross-sectional regressions. Size has a negative and significant premium, and book-to-market has a positive and significant premium. Notice that the use of the characteristic-based three factor model can explain about 35% of the cross-sectional variation of the equity premium. The high goodness-of-fit

Table 8 : The Cross-Section of Expected Equity Risk Premium

Panel A of this table reports the Fama-MacBeth cross-sectional regressions of firm-level equity risk premium on the market beta (β_{MKT}), SMB beta (β_{SMB}), HML beta (β_{HML}), and WML beta (β_{WML}), separately and jointly. These factor loadings are estimated via 60-month rolling regressions of the realized equity excess return on the Fama-French three factors and WML. Panel B reports the cross-sectional regressions of firm-level equity risk premium on the market beta, size, book-to-market, and past 12-month returns, separately and jointly. The Fama-MacBeth t-statistics reported in parentheses are adjusted for autocorrelations of up to 12 order in the slope coefficients using the method of Pontiff (1996). Regression coefficients are in percent.

Panel A: Covariance-Based Tests				Panel B: Characteristic-Based Tests		
	Model 1	Model 2	Model 3		Model 4	Model 5
Intercept	6.893	6.701	6.803	Intercept	32.295	32.301
	(12.53)	(12.96)	(13.19)		(21.37)	(21.63)
β_{MKT}	2.867	1.159	1.005	β_{MKT}	0.634	0.509
	(7.50)	(3.01)	(2.41)		(2.12)	(1.71)
β_{SMB}		5.398	5.668	log(ME)	-4.057	-3.905
		(16.79)	(17.02)		(-22.66)	(-22.39)
β_{HML}		3.125	3.566	B/M	5.727	5.310
		(8.86)	(10.02)		(10.62)	(10.63)
β_{WML}			-35.704	Past Returns		-6.047
			(-0.56)			(-9.36)
Average R^2	0.013	0.162	0.184	Average R^2	0.350	0.370

coefficients of Panel B are remarkable because our cross-sectional regressions are performed directly at the firm-level without portfolio formation.

The last column of Panel B (Model 5) reports the cross-sectional regressions of expected equity risk premium on market beta, size, book-to-market, and past 12-month returns. Size continues to have a negative and significant premium, and book-to-market continues to have a positive and significant premium. However, past realized 12-month returns load negatively on the expected equity risk premium with a coefficient of -6.04% per annum, and is highly significant (t-statistic of -9.36). With all three firm characteristics (size, book-to-market, and past returns) in the regression, the market beta has a risk premium of 0.51% per annum, which seems rather low, and is only significant at the 10% level (t-statistic of 1.71).

Implications

Our findings have potential important implications for what we know about the cross-section of equity returns. First, consistent with Fama and French (1993, 1996), the evidence we present indicates that their size and book-to-market factors are indeed risk factors that are priced *ex ante*. This differs from Brav et al. (2003) who find that the value premium does not exist in their data. Second, the market beta is significantly priced even after the size and book-to-market factors are jointly considered. This contrasts with the results in Fama and French (1992) where the market beta does not have any power in comparison with the size and book-to-market factors. Third, consistent with Brav et al., we also find that the momentum factor is not priced *ex ante* — momentum profits may be an empirical by-product of average, *ex post* return measures.

3.4 Robustness

Our results on expected returns are based on the specification of Eq. (8), which takes $\partial S_{it}/\partial B_{it}$ to be a linear function of the leverage ratio, B_{it}/S_{it}. This specification is admittedly simple, but provides for a natural benchmark. To examine whether our findings could be somehow explained by our model parsimony (misspecification), we conduct checks that use alternative specifications for $\partial S_{it}/\partial B_{it}$. In doing so, we revisit all of our paper's results on equity pricing.

Merton (1974) implies that $\frac{\partial S}{\partial B} = \frac{N(d_1)}{1-N(d_1)}$, where $N(\cdot)$ is the cumulative distribution function of a normal variate, $d_1 \equiv \frac{\log(F/B)+\left(r+\frac{1}{2}\sigma^2\right)T}{\sigma\sqrt{T}}$, and σ is the volatility of the firm's asset return. This model suggests that $\partial S_{it}/\partial B_{it}$ should also be a function of firm volatility and the risk free rate r_t, in addition to the leverage ratio. Following Merton (1974), we first redesign the estimate of $\partial S_{it}/\partial B_{it}$ as follows (with coefficient estimates and their t-statistics):

$$\partial S_{it}/\partial B_{it} = \underset{(29.7)}{23.08} + \underset{(19.06)}{2.07}\,(B_{it}/S_{it}) + \underset{(2.24)}{12.56}\,\sigma_{it} - \underset{(15.27)}{1.39}\,r_t + \varepsilon_{it}, \tag{10}$$

where we use equity volatility, σ_{it}, as opposed to firm volatility. Alternatively, as a less model-dependent but more inclusive specification, we also model $\partial S_{it}/\partial B_{it}$ as:

$$\partial S_{it}/\partial B_{it} = \underset{(-6.01)}{-11.80} - \underset{(1.04)}{0.15}\,(B_{it}/S_{it}) + \underset{(7.96)}{47.35}\,\sigma_{it} - \underset{(6.38)}{0.62}\,r_t + \underset{(20.49)}{3.99}\,\log(\text{ME}) - \underset{(10.56)}{6.26}\,(\text{B/M}) + \underset{(6.18)}{5.55}\,r_{it}^{12} + \varepsilon_{it}, \quad (11)$$

essentially augmenting Eq. (10) with the natural log of market value, log(ME), and the past 12-month equity return, r_{it}^{12}.

Table 9 reports the properties of the common equity factors that are constructed under the two alternative proxies for $\partial S_{it}/\partial B_{it}$ (i.e., Eqs. (10) and (11)). Overall, the results are similar to those of Table 5. We find significantly positive premia for the market and value factors, and a significantly negative momentum premium. The size premium is significantly positive with specification (10), but is insignificant with (11), albeit still positive.

Table 10 reports the cross correlations of the expected equity factor returns (constructed with the alternative proxy for $\partial S_{it}/\partial B_{it}$) with the cyclical component of the real industrial production index. In general, the patterns are unchanged from Table 6, i.e., the expected market, size, and value premia are countercyclical and the expected momentum return is procyclical. However, Panel B of Table 10 shows that when $\partial S_{it}/\partial B_{it}$ is modelled as Eq. (11), the expected size and momentum returns are both acyclical, as their cross correlations with output are not significantly different from zero.

Table 11 reports the cross-sectional regressions of expected equity returns on loadings of Fama-French factors and firm characteristics using the two alternative specifications of $\partial S_{it}/\partial B_{it}$ from Eqs. (10) and (11). The results are very similar to those in Table 8. The market beta is always positive and significant even after we control for the other factor loadings as well as for firm characteristics. The loadings on SMB and HML are significant and positive in all of the regressions.

3.5 The Cross-Section of Yield Spreads

This subsection examines the empirical relation between the cross-sectional variation of yield spreads, common equity risk factors, and firm characteristics. While it is standard to assume that the market equity risk premium is a linear function of the yield spread at the aggregate

Table 9 : Expected Returns of Common Equity Factors: Robustness Checks

This table reports summary statistics of expected returns of common equity factors, including market excess return (MKT), SMB, HML, and WML (the momentum factor). We report mean, standard deviation (std), min, max, and autocorrelations of orders 1 (ρ_1), 2 (ρ_2), 6 (ρ_6), and 12 (ρ_{12}), the results of market regressions including the intercepts (α) and the slopes (β) as well as their Newey-West t-statistics, as well as the correlation matrix for these four factors. The numbers of mean, min, max, and α are in annualized percent. All cross-correlations are significant at the 1-percent test level or lower. All the t-statistics are adjusted for heteroscedasticity and autocorrelations of up to 12 lags. Expected returns are constructed using two alternative specifications of $\partial S_{it}/\partial B_{it}$.

Panel A: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \varepsilon_{it}$								
	mean	std	min	max	ρ_1	ρ_2	ρ_6	ρ_{12}
MKT	4.369	0.150	0.535	12.348	0.876	0.794	0.490	0.329
SMB	4.381	0.212	-6.376	20.021	0.816	0.649	0.328	-0.052
HML	7.764	0.324	-9.283	22.943	0.895	0.830	0.590	0.439
WML	-1.472	0.307	-19.746	22.007	0.820	0.701	0.347	0.146
	Cross-Correlations				Market Regressions			
	MKT	SMB	HML	WML	α	t_α	β	t_β
MKT	1	0.354	0.558	-0.308	na	na	na	na
SMB		1	-0.110	-0.263	2.263	5.589	0.483	6.079
HML			1	-0.164	2.671	4.877	1.163	10.811
WML				1	1.268	2.175	-0.627	-5.423

Panel B: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \alpha_4 \log(\text{ME}) + \alpha_5(\text{B/M}) + \alpha_6 r_{it}^{12} + \varepsilon_{it}$								
	mean	std	min	max	ρ_1	ρ_2	ρ_6	ρ_{12}
MKT	5.117	0.196	1.731	15.557	0.940	0.896	0.693	0.503
SMB	0.101	0.155	-8.123	8.629	0.762	0.603	0.332	0.121
HML	7.650	0.364	-2.556	26.269	0.903	0.853	0.740	0.560
WML	-0.723	0.297	-18.849	12.004	0.771	0.705	0.298	0.116
	Cross-Correlations				Market Regressions			
	MKT	SMB	HML	WML	α	t_α	β	t_β
MKT	1	-0.499	0.689	-0.346	na	na	na	na
SMB		1	-0.475	0.077	2.179	8.219	-0.389	-9.234
HML			1	-0.243	0.926	1.781	1.259	15.234
WML				1	1.949	3.791	-0.522	-6.174

Table 10 : Cross Correlations with Cyclical Component of Industrial Production: Robustness Checks

This table reports the cross correlations of expected returns of equity factors with the cyclical component of the industrial production index, corr(r_t, y_{t+k}), for different leads and lags, k. The cyclical component of the real industrial production index (obtained from FRED) is estimated by passing the raw series through the Hodrick and Prescott (1997) filter. Expected returns are constructed using two alternative specifications of $\partial S_{it}/\partial B_{it}$. p-values of the cross-correlations are reported in parentheses.

Panel A: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \varepsilon_{it}$													
k	-24	-12	-6	-3	-2	-1	0	1	2	3	6	12	24
MKT	0.132	0.062	-0.091	-0.224	-0.299	-0.361	-0.372	-0.372	-0.344	-0.300	-0.139	0.135	0.234
	(0.03)	(0.31)	(0.13)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.02)	(0.03)	(0.00)
SMB	-0.117	-0.003	-0.215	-0.341	-0.362	-0.373	-0.392	-0.403	-0.380	-0.336	-0.161	0.236	0.270
	(0.06)	(0.97)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
HML	0.072	-0.032	-0.083	-0.249	-0.300	-0.319	-0.322	-0.307	-0.274	-0.234	-0.109	0.053	0.212
	(0.25)	(0.61)	(0.18)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.08)	(0.41)	(0.00)
WML	-0.348	-0.200	-0.071	0.120	0.201	0.255	0.259	0.269	0.251	0.233	0.170	0.001	-0.021
	(0.00)	(0.00)	(0.24)	(0.04)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.98)	(0.73)
Panel B: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \alpha_4 \log(\text{ME}) + \alpha_5(\text{B/M}) + \alpha_6 r_{it}^{12} + \varepsilon_{it}$													
k	-24	-12	-6	-3	-2	-1	0	1	2	3	6	12	24
MKT	0.310	0.116	-0.080	-0.181	-0.229	-0.270	-0.283	-0.288	-0.274	-0.245	-0.144	0.009	0.230
	(0.00)	(0.06)	(0.19)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.02)	(0.89)	(0.00)
SMB	-0.212	-0.094	0.055	0.068	0.093	0.110	0.095	0.064	0.031	0.010	-0.029	0.011	0.022
	(0.00)	(0.13)	(0.38)	(0.28)	(0.14)	(0.08)	(0.13)	(0.31)	(0.62)	(0.87)	(0.65)	(0.87)	(0.74)
HML	0.168	0.018	-0.171	-0.214	-0.234	-0.240	-0.239	-0.237	-0.236	-0.232	-0.181	0.007	0.120
	(0.01)	(0.77)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.91)	(0.07)
WML	-0.354	0.000	-0.129	-0.046	0.003	0.007	0.011	0.019	0.005	0.002	0.019	0.098	0.025
	(0.00)	(1.00)	(0.03)	(0.44)	(0.97)	(0.91)	(0.86)	(0.75)	(0.93)	(0.98)	(0.75)	(0.11)	(0.69)

level (see, e.g., Ferson and Harvey (1991) and Jagannathan and Wang (1996)), to the best of our knowledge, the literature lacks a direct assessment of the degree to which systematic risk drives (firm-level) cross-sectional variation of yield spreads.[14] Contradicting this practice, many studies in the default risk literature assume a zero risk premium, implying that the risk in the yield spread is purely idiosyncratic (e.g., Bodie et al. (1993), Fons (1994), and Cumby and Evans (1995)). Sorting out which approach is more appropriate seems important.

Table 12 reports the results from cross-sectional regressions of yield spreads on common

[14]Elton et al. (2001) look at whether the *residual* of yield spreads can be explained by the Fama-French factors, but do not study whether the gross yield spread can be explained by those factors. Importantly, those authors restrict their analysis to investment grade bonds, excluding the types of risky firms that are necessary for studying equity returns.

Table 11 : The Cross-Section of Expected Equity Risk Premium: Robustness Checks

This table reports the Fama-MacBeth cross-sectional regressions of firm-level equity risk premium on the market beta (β_{MKT}), SMB beta (β_{SMB}), HML beta (β_{HML}), and WML beta (β_{WML}), as well as size, book-to-market, and past 12-month returns, separately and jointly. These factor loadings are estimated via 60-month rolling regressions of the realized equity excess return on the Fama-French three factors and WML. The Fama-MacBeth t-statistics reported in parentheses are adjusted for autocorrelations of up to 12 order in the slope coefficients using the method of Pontiff (1996). All the regression coefficients are in percent. Expected returns are constructed using two alternative specifications of $\partial S_{it}/\partial B_{it}$.

Panel A: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \varepsilon_{it}$						
	Covariance-Based Tests			Characteristic-Based Tests		
	Model 1	Model 2	Model 3		Model 4	Model 5
Intercept	3.772	3.709	4.024	Intercept	14.093	12.752
	(6.18)	(5.43)	(5.48)		(6.32)	(5.62)
β_{MKT}	2.951	2.047	1.751	β_{MKT}	2.195	2.054
	(4.79)	(4.76)	(4.01)		(4.83)	(4.29)
β_{SMB}		3.429	3.581	log(ME)	-1.911	-1.739
		(11.76)	(11.68)		(-8.26)	(-7.49)
β_{HML}		1.316	1.728	B/M	5.776	5.794
		(3.08)	(3.59)		(4.87)	(5.00)
β_{WML}			13.223	Past Returns		-0.109
			(0.18)			(-0.10)
Average R^2	0.031	0.186	0.218	Average R^2	0.398	0.415

Panel B: $\partial S_{it}/\partial B_{it} = \alpha_0 + \alpha_1(B_{it}/S_{it}) + \alpha_2\sigma_{it} + \alpha_3 r_t + \alpha_4 \log(\text{ME}) + \alpha_5(\text{B/M}) + \alpha_6 r_{it}^{12} + \varepsilon_{it}$						
	Covariance-Based Tests			Characteristic-Based Tests		
	Model 1	Model 2	Model 3		Model 4	Model 5
Intercept	4.345	4.629	4.870	Intercept	10.219	8.526
	(6.94)	(6.94)	(6.76)		(5.30)	(4.52)
β_{MKT}	2.478	1.419	1.164	β_{MKT}	1.776	1.679
	(3.52)	(2.92)	(2.49)		(3.53)	(3.26)
β_{SMB}		2.760	2.894	log(ME)	-1.419	-1.230
		(9.39)	(9.33)		(-6.91)	(-5.89)
β_{HML}		1.574	1.899	B/M	7.123	7.249
		(4.08)	(4.18)		(6.01)	(6.48)
β_{WML}			-55.636	Past Returns		0.675
			(-0.88)			(0.68)
Average R^2	0.022	0.112	0.135	Average R^2	0.192	0.222

equity factors and firm characteristics. Model 1 in Panel A indicates that when the market factor is used as the regression's only explanatory variable its slope attracts a positive, significant coefficient of 0.77. Model 2 suggests that the addition of the SMB and HML betas to the regression does not eliminate the significance of the market beta. Together, the Fama-French three factor loadings can explain on average 29% of the cross-sectional variation of yield spreads. The third column in Panel A (Model 3) shows that the momentum factor does not play a significant role in explaining the cross-section of yield spreads. In Model 4, we add bond rating, a variable that (presumably) should summarize the default risk. Its slope is positive and highly significant (as expected), and the average R^2 goes up to 56%. Yet, the Fama-French three factors remain positive and significant, while the slope of the momentum factor becomes negative and insignificant. The addition of equity volatility, leverage ratio, and volatility increases the R^2 further to 66%, as reported in Model 5.

In Panel B of Table 12, we replace the SMB, HML, and WML factor loadings with log size, book-to-market, and past 12-month returns. The patterns are similar to those from Panel A, but the estimates are generally more significant. The three factor model (Model 6) can explain on average 49% of the cross-sectional variation of the yield spreads. The addition of all variables increases the total R^2 to 68% (Model 9). Remarkably, out of the portion of the yield spread variation that can be explained, about 72% is related to the Fama-French three risk factors. These results essentially imply that the yield spread is highly correlated with systematic risk, supporting our notion that there exists a strong link between equity and bond markets.

4 Conclusion

We construct measures of expected returns using bond data and examine the relation between risk and expected equity returns. Our tests do not assume that the average realized returns is an unbiased proxy for the *ex ante* expected return. We find that: (i) the market beta plays a significant role in the cross-section of expected returns, and its role persists even after size and book-to-market are controlled for; (ii) the risk premia associated with size and book-to-market are positive, significant, and countercyclical; (iii) there is little evidence on

Table 12 : The Cross-Section of Bond Yield Spreads

Panel A of this table reports the Fama-MacBeth cross-sectional regressions of firm-level bond yield spreads on the market beta (β_{MKT}), SMB beta (β_{SMB}), HML beta (β_{HML}), and WML beta (β_{WML}) as well as bond ratings, stock return volatility, leverage, and firm-specific average maturity of bonds. These factor loadings are estimated via 60-month rolling regressions of the realized equity excess return on the Fama-French three factors and WML. Panel B replicates the analysis in Panel A, but replacing the loadings on SMB, HML, and WML by size, book-to-market, and past 12-month, respectively. The Fama-MacBeth t-statistics reported in parentheses are adjusted for autocorrelations of up to 12 order in the slope coefficients using the method of Pontiff (1996). Regression coefficients are in percent.

Panel A: Covariance-Based Tests						Panel B: Characteristic-Based Tests				
	Model 1	Model 2	Model 3	Model 4	Model 5		Model 6	Model 7	Model 8	Model 9
Intercept	0.971	0.916	0.886	-1.408	-1.917	Intercept	6.079	5.993	1.134	-0.014
	(8.49)	(8.92)	(8.74)	(-9.97)	(-15.83)		(32.10)	(30.11)	(6.28)	(-0.08)
β_{MKT}	0.774	0.534	0.559	0.126	-0.021	β_{MKT}	0.384	0.336	0.076	-0.020
	(7.10)	(6.25)	(6.88)	(2.24)	(-0.47)		(6.73)	(6.59)	(1.71)	(-0.58)
β_{SMB}		1.028	1.085	0.262	0.181	log(ME)	-0.697	-0.677	-0.268	-0.198
		(24.66)	(23.94)	(7.90)	(5.47)		(-31.50)	(-32.14)	(-17.31)	(-14.63)
β_{HML}		0.123	0.159	0.072	0.035	B/M	0.302	0.286	0.232	0.202
		(2.45)	(2.91)	(2.00)	(1.20)		(3.96)	(3.69)	(3.09)	(3.01)
β_{WML}			4.886	-9.584	-3.679	Past Returns		-0.256	-0.549	-0.277
			(0.45)	(-1.25)	(-0.77)			(-2.43)	(-5.88)	(-4.54)
Rating				0.823	0.584	Rating			0.647	0.521
				(20.43)	(24.85)				(26.25)	(28.98)
Volatility					10.998	Volatility				9.574
					(9.45)					(11.08)
Leverage					1.387	Leverage				0.841
					(13.98)					(9.77)
Maturity					0.008	Maturity				0.008
					(3.72)					(4.00)
Average R^2	0.028	0.290	0.326	0.563	0.656	Average R^2	0.474	0.489	0.611	0.679

positive momentum profits. We also find that the common equity risk factors are the main determinants of the cross-sectional variation of the yield spreads.

While we do not claim that our expected return measure should dominate any other, we believe that re-examining some of the inferences of the asset pricing literature from the last 20 years (most of which use *ex post* returns) with an *ex ante* return measure constructed from first principles to be a valid experiment. Since our proposed measure captures information that — both on theoretical and empirical grounds — is shown to be imperfectly correlated with that of *ex post* average returns, we think that experiments like ours may shed some new light into how one should understand the pricing of risky securities in the financial markets.

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A Proofs

Proof of Proposition 1: Let the uncertainty be represented by a N-dimensional Brownian motion $W_t = (w_{1t}, w_{2t}, \ldots, w_{Nt})'$. There are M firms, $F_1, F_2, \ldots, F_M$ whose asset values obey the following well-defined stochastic processes:

$$\frac{dF_{it}}{F_{it}} = \mu_{it}\, dt + \sigma'_{it}\, dW_t \tag{A1}$$

where $i = 1, \ldots, M$ and $\sigma_{it} \equiv (\sigma^i_{1t}, \sigma^i_{2t}, \ldots, \sigma^i_{Nt})'$. Following Merton (1974), we assume that all firms are levered with predetermined debt. For firm i, both the equity price S_{it} and debt value B_{it} will depend on the underlying asset value F_{it}.

Assume that the state price density process, Λ_t, is given by:

$$d\Lambda_t = \mu_{\Lambda t} dt + \sigma'_{\Lambda t}\, dW_t. \tag{A2}$$

It follows that firm i's expected excess return on its asset is given by:

$$R^i_{Ft} - r_t = -\text{Cov}_t\left(\frac{dF_{it}}{F_{it}}, \frac{d\Lambda_t}{\Lambda_t}\right). \tag{A3}$$

where $R^i_{Ft} \equiv \text{E}_t\left[\frac{dF_{it}}{F_{it}}\right]/dt$ is the expected return on the assets, $\text{Cov}_t\left(\frac{dF_{it}}{F_{it}}, \frac{d\Lambda_t}{\Lambda_t}\right)$ denotes the instantaneous conditional covariance (normalized by dt), and $r_t \equiv -\mu_{\Lambda t}/\Lambda_t$ is the real interest rate.

As equity S_{it} and debt B_{it} are contingent claims written on the same underlying productive asset, an application of Itô's lemma yields the risk premia for these two securities:

$$R^i_{St} - r_t \;=\; -\left(\frac{\partial S_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{S_{it}}\right)\text{Cov}_t\left(\frac{dF_{it}}{F_{it}}, \frac{d\Lambda_t}{\Lambda_t}\right) \tag{A4}$$

$$R^i_{Bt} - r_t \;=\; -\left(\frac{\partial B_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{B_{it}}\right)\text{Cov}_t\left(\frac{dF_{it}}{F_{it}}, \frac{d\Lambda_t}{\Lambda_t}\right) \tag{A5}$$

Both equity risk and debt risk premia are determined by the systematic component of the underlying asset, $\text{Cov}_t\left(\frac{dF_{it}}{F_{it}}, \frac{d\Lambda_t}{\Lambda_t}\right)$. Equation (1) in Proposition 1 follows by taking the ratio of (A4) and (A5).
■

Proof of Proposition 2: Similar to Jarrow (1978), we start with the bond yield equation:

$$B_{it} = \sum_{j=1}^{n} C_i e^{-Y_{it}T_j} + K_i e^{-Y_{it}T_n}, \tag{A6}$$

where C_i is the coupon payment of the bond, n is the number of remaining coupons, Y_{it}

is the bond's yield to maturity, T_j, $j = 1, \dots, n$ are length of time period for each coupon payment, and K_i is the face value of debt.

As the bond yield, Y_{it}, is the only time-varying variable on the right-hand side of (A6), by Itô's Lemma we can write bond risk premium as a function of the bond yield and other observable bond characteristics:

$$\frac{\mathrm{E}_t\left[dB_{it}\right]}{B_{it}} = \mathrm{E}_t\left[\frac{\partial B_{it}}{\partial t}\frac{dt}{B_{it}} + \frac{\partial B_{it}}{\partial Y_{it}}\frac{dY_{it}}{B_{it}} + \frac{1}{2}\frac{\partial^2 B_{it}}{\partial Y_{it}^2}\frac{1}{B_{it}}\left(dY_{it}\right)^2\right] \tag{A7}$$

where

$$\frac{\partial B_{it}}{\partial t} = Y_{it}B_{it} \tag{A8}$$

$$\frac{\partial B_{it}}{\partial Y_{it}} = -H_{it}B_{it} \quad \text{with} \quad H_{it} = \sum_{j=1}^{n}\frac{T_j C_i e^{-Y_{it}T_j}}{B_{it}} + \frac{T_n K_i e^{-Y_{it}T_n}}{B_{it}} \tag{A9}$$

$$\frac{\partial^2 B_{it}}{\partial Y_{it}^2} = G_{it}D_{it} \quad \text{with} \quad G_{it} \equiv \sum_{j=1}^{n}\frac{T_j^2 C_i e^{-Y_{it}T_j}}{B_{it}} + \frac{T_n^2 K_i e^{-Y_{it}T_n}}{B_{it}} \tag{A10}$$

and H_{it} and G_{it} are modified duration and convexity, respectively. (A7) thus becomes:

$$R_{Bt}^i - r_t = \mathrm{E}_t\left[\frac{dB_{it}}{B_{it}}\right]/dt - r_t = (Y_{it} - r_t) - H_{it}\frac{\mathrm{E}_t\left[dY_{it}\right]}{dt} + \frac{1}{2}G_{it}\frac{\left(dY_{it}\right)^2}{dt} \tag{A11}$$

■

Proof of Proposition 3: The proposition follows by combining (2) with:

$$\begin{aligned}
\mathrm{E}_t[dY_{it}] &= \pi_{it}\mathrm{E}_t[dY_{it}^-] + (1-\pi_{it})\mathrm{E}_t[dY_{it}^+] \\
(dY_{it})^2 &= \pi_{it}(dY_{it}^-)^2 + (1-\pi_{it})(dY_{it}^+)^2
\end{aligned}$$

■